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Abstract

This paper explores the determinants of the proportion of the working age male population claiming incapacity benefits (IB), across the eleven British Government Office Regions, for the period 1998-2006. Three different approaches are adopted to modelling register dynamics: first treating IB stocks as if they were trend-stationary, albeit with persistence, and estimating reduced form models for their logs; second treating IB stocks as if they were non-stationary and examining their long run determinants plus short run equilibrium reversion properties; third focusing on the determinants of gross inflows and outflows that together drive IB stocks. Given the nature of the data no approach is ideal, yet the models provide reasonably robust evidence that labour market changes – specifically falling unemployment rates and rising real earnings – have contributed to falling male IB stocks over the period.

1. Introduction

The number of people of working age claiming income replacement incapacity benefits (henceforth IB[1]) in Britain currently stands at 2.7 million. This figure has grown by over 300 percent in 30 years, although the growth rate has slowed considerably in recent years. Similar growth has been experienced by the US (e.g. see Bound and Burkhauser, 1999; Autor and Duggan, 2006) and by many other OECD countries (e.g. see Bound and Burkhauser, 1999; Prinz, 2003). In both the UK and US the headline figures are split unequally by gender, with men making up around three fifths of the totals.

McVicar (2008) suggests that the causes of the phenomenal growth in British IB rolls are still poorly understood, or at least poorly *quantified*. This is despite a number of informative descriptive and qualitative studies in recent years (e.g. see Huddleston, 2000; Walker and Howard, 2000; Alcock et al., 2003; Bell and Smith, 2004). McVicar's argument is that estimation of properly specified multivariate models is likely to be necessary – although perhaps not sufficient – in order to get a clearer idea of the relative roles played by disability prevalence, benefit characteristics and the labour market in explaining the growth of IB rolls. Few papers have attempted such an exercise using British data. The US literature is more convincing in this respect, but institutional and other differences mean US findings may not be directly exportable.

[1] Here we follow the usual convention in the British incapacity benefit literature of using the term 'incapacity benefits' to denote all income replacement incapacity benefits, i.e. Incapacity Benefit itself, but also Severe Disablement Allowance and Credits Only cases. We use the abbreviation 'IB' as a shorthand for all these incapacity benefits, not just for Incapacity Benefit, which itself is commonly abbreviated to 'IB'.

The growth in IB rolls has not been uniform across space, and within some countries this has led to the emergence of significant regional differences in the proportion of the regional working age populations claiming such benefits. In the US, for example, the southern states generally have a higher proportion of the working age population receiving Disability Insurance (DI) than other states (e.g. see Rupp and Stapleton, 1998; McVicar, 2006). Regional differences in Britain are particularly pronounced, with Wales and the North having considerably higher IB rolls than the South (e.g. see Fothergill and Wilson, 2007; McVicar, 2006). The causes of the differential growth rates that lie behind the current regional variation in British IB rolls – as in the case of aggregate national growth in IB rolls – are also, as yet, poorly understood. Again, although we may have a reasonable qualitative understanding of the factors behind regional variation, few studies have attempted to quantify the relative roles of the various factors believed to drive regional IB rolls, and those that have done so suffer from a variety of problems. As with explaining aggregate growth, there are likely to be limits in terms of how much we can learn about Britain from US data.

In this paper we use quarterly administrative and Labour Force Survey (LFS) data for the eleven British Government Office Regions (GORs) over the period 1998-2006 to estimate reduced form models of male regional IB rolls. The model treats IB rolls as being determined by a combination of disability prevalence, labour market and demographic factors, and unobserved time and region fixed effects. Different approaches are explored to the specification of the dynamics of IB rolls, including lagged dependent variable (LDV) models, error correction models (ECMs) and models that focus explicitly on the flows onto and off IB. Existing quantitative studies of British IB rolls have tended to skirt over issues of dynamics and/or have omitted either disability prevalence or labour market factors. To a lesser extent, this is also true of existing US studies of DI rolls using aggregate level data.

Much of the existing quantitative literature on IB rolls examines either only male claimants or all claimants without distinguishing between males and females. Our view is that the determinants of male IB claims and female IB claims may differ, perhaps because labour force participation decisions are more complex for many women than for most men. Recent trends in IB rolls are consistent with such differences, e.g. with the number of male IB claimants falling by ten percent over our sample period but the number of female IB claimants rising by 14 percent. Where studies have examined male and female claimants separately, some have found shared determinants with similar magnitudes and others have not. To keep this paper a manageable size, we focus on male claimants. Given the growth in the share of IB claimants that are female, however, we also provide a link to an online appendix with a brief discussion focussed on female claimants, and presenting results from estimating the same models on female data as are presented here for males.[2]

In the following section the quantitative literature on growth and spatial differences in IB rolls is reviewed. Our focus is on British data, but we also draw selectively on the more substantial US literature. Section 3 briefly describes the data. Section 4 presents and discusses estimation results when we treat the IB rolls as (trend) stationary. Section 5 presents and discusses first-difference and ECM estimates when we treat IB rolls as if they are non-stationary. Section 6 presents and

[2] www.erini.ac.uk/sandproject.

discusses estimates of the factors explicitly determining flows onto and off the IB register and discusses how these in turn affect the stock of IB claimants. Section 7 concludes.

2. Existing Quantitative Literature on IB Rolls

The theory of labour supply is most frequently called upon in the empirical economics literature on IB rolls, whether explicitly, e.g. as in Kreider and Riphahn (2000) and Autor and Duggan (2003), or implicitly, e.g. as in Molho (1989, 1991) and Disney and Webb (1991). Specifically, we can imagine a model where individuals seek to maximize the present value of their expected utility by choosing whether to participate in the labour market or claim IB in each period. This choice will be influenced by their disability status, their probability of getting a job given participation, the probability of receiving disability benefit given a claim, the value of disability benefits relative to the expected wage, and unobserved heterogeneity. There may also be some degree of state dependence (inertia within a given state). However, such models, whether at individual or aggregate level, and whether structural or reduced form, have rarely been estimated on British data.

First, there is very little in the way of quantitative time series analysis of the determinants of the British IB rolls. Disney and Webb (1991) explore estimation of a reduced form model for the number of Invalidity Benefit (IVB)[3] claimants – omitting any measure of disability prevalence – using annual data from 1962 to 1989, but argue that time-series specifications are problematical given the non-stationarity of the main series and the apparent lack of a cointegrating vector between them (the IVB register, the unemployment rate and the real value of IVB). Nevertheless,

[3] IVB was the predecessor to Incapacity Benefit.

they report estimates from a regression in levels, treating the main series as stationary, suggesting a positive relationship between IVB rolls and unemployment (elasticity of 0.3 to 0.8) but a counterintuitive negative relationship between IVB rolls and the real value of IVB. Although not explicitly a study of the determinants of IB rolls, Bell and Smith (2004) estimate a single equation reduced form model for the male inactivity rate 1984-2001, using the 1995 reforms to help identify the labour supply effects of the generosity of IB. They find the real level of IB to have a positive and significant impact on inactivity (elasticity of 0.26), but do not control for disability prevalence, labour market factors or other benefit characteristics beyond a linear trend and time dummies. There appears to be a similar lack of published, quantitative, aggregate, time series analysis of US DI rolls (see Bound and Burkhauser, 1999).

There are two existing British regional-level panel studies of IB rolls. Disney and Webb (1991) estimate the proportion of a region's eligible population claiming IVB using three time points for twelve UK regions between 1980 and 1988. They find intuitively signed and statistically significant impacts of unemployment rates (elasticity of 0.1), real IVB rates (elasticity of 0.3) and population age structure on regional IVB rolls. By once again omitting any measure of disability prevalence, however, they are unable to explore the extent to which variation in prevalence contributes to variation in IVB rolls and, to the extent that such variation is not controlled for by the fixed effects, their estimates may suffer from omitted variable bias. The study is also now somewhat dated, predating as it does the period of most rapid increase in IB rolls (1988-1995), the recent sustained period of falling unemployment across Britain (1993-2005), and various significant policy changes to incapacity and other benefits. More recently, Faggio and Nickell (2005) use various survey data for 1972-2002 to estimate a fixed effects model for the proportion of prime age men that are inactive. They do include the proportion of the relevant regional population reporting a disability, which they treat as exogenous, but find it plays no role in explaining inactivity rates given the (region and time) fixed effects. They also find counter-intuitively signed relationships between inactivity and the real IB weekly rate (-), and between inactivity and vacancy rates (+) and employment growth (+), again given the fixed effects. Regional average wages have a correctly signed coefficient with elasticity of -1.

Two recent US studies examine the determinants of state-level DI rolls using panel data. To study the role of economic incentives in determining DI rolls, Black et al. (2002) regress the change in county-level DI rolls on change in average earnings, where this is instrumented by exogenous changes in the value of coal reserves during the coal boom and bust of the 1970s/80s, together with some controls, across four US states. The controls do not contain any measure of disability prevalence or any other measure of the state of the labour market. Given their focus, the model structure is designed to deal with the potential joint determination of average earnings and DI rolls, with both likely to be influenced by disability prevalence, which is unobserved. So, although Black et al. plausibly address the problem of omitted variable bias with regard to disability prevalence, its omission leaves them unable to explore the extent to which its variation contributes directly to variation in DI rolls. This aside, Black et al. (2002) provide convincing evidence that economic incentives are important determinants of DI rolls, with county DI rolls negatively and significantly related to average earnings (elasticity of around -0.4). Autor and Duggan (2003) similarly estimate a simple model for the change in the proportion of working age adults receiving DI at the US state level. The only included disability prevalence measure is the state mortality rate (the level rather than the change), which is shown to be uncorrelated with changes in DI rolls. It cannot be ruled out, however, that the lack of a relationship between mortality rates and DI is driven by the fact that mortality rates are a very imperfect proxy for disability prevalence (see Bound et al., 1991). Autor and Duggan include a measure of labour demand – constructed from state industry-mix weighted changes in national sectoral employment in order to ensure exogeneity – but this too is statistically insignificant. DI replacement rates, however, are positive and significant with an elasticity of 0.5.

A number of US studies examine the determinants of state-level DI *applications* using panel data. Parsons (1991) estimates an equation for DI applications in first differences across 42 US states over the period 1977-1980. He finds an elasticity of -0.5 with respect to initial denial rates (interpretable as a proxy for screening stringency), yet includes no other controls. First differencing will only remove time-invariant heterogeneity across states, so Parsons implicitly assumes no change in disability incidence, no change in labour demand, and no change in replacement rates over the period. We have already seen, at least in the case of the latter two, that these omitted factors are likely to matter in the US. Stapleton et al. (1998) estimate an equation for DI application rates across US states over the period 1988-1992. They also find a significant impact of denial rates, although with an elasticity half that found by Parsons (1991). They also include a number of other controls, including the unemployment rate (+), demographic controls and a measure of the incidence of AIDS/HIV. Autor and Duggan (2003) use a variety of specifications to show the number of DI applications – in contrast to the DI stock – does depend on their measure of labour demand. None of these specifications include any measure of disability prevalence.

Turning to aggregate level cross section studies, Nolan and Fitzroy (2003) use British Local Authority (LA) level LFS data for three years (1999-2001), as three separate cross sections, in order to estimate what they describe as preliminary regressions for the proportion of the LA working age population claiming IB. Their equation includes hospital visits (+) and mortality rates (+) proxying for the underlying population health, some socio-economic indicators, but no labour market controls and no measure of replacement rates. McVicar (2007) uses 2003 British LA level LFS data to estimate a cross section model for IB rolls allowing for self-reported disability prevalence, average earnings and local unemployment rates to be endogenously determined. IB rolls for both men and women are shown to vary positively with disability incidence and local unemployment rates, and negatively with average earnings. The elasticities are similar for males and females – perhaps surprisingly so given the recent divergence in male and female IB rolls – at around +1 for disability prevalence, +1/2 for unemployment rates and around -1 for median earnings.

Three British studies have used cross section data at the individual level to examine the determinants of claiming disability benefits, but they are somewhat dated, all using data from the 1980s. Disney and Webb (1991) use Family Expenditure Survey data separately for 1980, 1984 and 1988 to estimate a reduced form single equation model for IVB claims, with local unemployment rates taking an intuitive (positive) sign, although they are insignificant in 1988. They include no disability measure, however, beyond a dummy for smoking, which is not treated explicitly as an instrument. Molho (1989) and Molho (1991) estimate a single equation model for the probability of entry to IVB for males and females respectively. The estimated equation includes self-reported health status (treated as exogenous), local unemployment rates, weekly rate of IVB, together with further controls including some retrospective employment information. He finds health, demographic factors, benefit rates and past pay to be significant, with intuitive signs, for both men and women. Local unemployment rates are significant (with positive sign) for women but not men, although individual history of unemployment is significant for men. More recently, Jones et al. (2006) uses individual level data from the 2001 LFS to examine the determinants of labour force participation separately for the disabled and non-disabled and separately by gender. They find qualifications to be a key determinant of participation probabilities for the disabled, with significant local area dummies for women but not for men.

For the US, Kreider (1998) uses cross section survey data (with some additional retrospective information) from 1978 to examine the probability of a DI application in a structural model that focuses on the role of uncertainty over expected earnings in influencing the probability to apply for DI. The model allows a constructed measure of disability – based on various self-reported information but allowing for potential justification bias – and labour market factors to influence the application probability both directly and indirectly through uncertainty over earnings. Kreider's estimates suggest that disability (+), the benefit rate (+), the proportion of applications that are successful (+), and the unemployment rate (+) are all significant in the DI applications equation. Kreider and Riphahn (2000) use longitudinal data over the period 1986-93 to estimate a structural system of equations including a DI application equation, using a constructed measure of disability like Kreider (1998) to account for potential justification bias in self-reported disability status. For both men and women, the probability of applying for DI depends on disability (+), amongst other things. Unemployment rates and application rejection rates are not significant, and the present value of disability benefits is only significant for women. There are no such studies using British longitudinal data.

3. Data Details and Descriptive Analysis

Quarterly data on the number of working age IB claimants, since 1998Q1, by region and gender, are publicly available from the Department for Work and Pensions (DWP) website. On average over the period 1998Q1-2006Q1, the proportion of the male working age population claiming IB ranges from a low of 4.7 percent in the Southeast of England to a high of 13.8 percent in Wales. Overall, as can be seen in Figure 1, the regional pattern of male IB rolls follows a north-south divide, with the 'north' interpreted broadly to include Wales. This regional pattern has remained broadly stable over the study period, despite some degree of convergence (larger falls in 'northern' IB rolls). The national male IB roll fell by ten percent between 1998 and 2006.

<Figure 1 around here>

Quarterly data on the number of inflows to IB and outflows from IB are just becoming publicly available, although these data are only available from 1999Q4 and are still described by the Office for National Statistics (ONS) as 'experimental' (and therefore may be subject to revision). Nevertheless the estimated flow data can provide a useful alternative approach to estimating the determinants of IB rolls based on stock data only. Briefly, these estimated flow data suggest low inflow and outflow rates for all regions over all quarters, i.e. low churn; a stable regional hierarchy for inflow rates following the expected north-south pattern, with a downward trend more pronounced for the 'north' than the 'south'; seasonal variation in the estimated outflow rates but again evidence suggesting a downward trend.

<Figures 2 and 3 around here>

Turning to the explanatory factors suggested in the literature, quarterly LFS data are available for a number of variables from NOMIS and the ESRC Data Archive. These data include – by region and age group – information on the number of men describing themselves as work-limiting or Disability Discrimination Act (DDA) disabled, information on economic activity, information on regional industry mix, and demographic information. These data are used to construct variables for the proportion of working age men reporting a disability, employment and unemployment rates by disability status and overall, a variable for the proportion of the working age population aged 50 years or over, and a variable for the proportion of the employed working in manufacturing (the latter two are not gender specific). Quarterly data on the median and tenth percentile weekly earnings of full-time male employees by region are available from the Annual Survey of Hours and Earnings (and its predecessor) again available from NOMIS and the ONS.

Summary statistics for these series, by region, are reported in Table 1. Notice that within-region variation is limited for most regions and most variables, partly because the sample period is short, but also because it coincides with a relatively stable period for IB rolls and the labour market. This will leave the region fixed effects to pick up a lot of the 'explanatory power' of the model, which is a fundamental constraint in such a regional panel exercise where it is difficult to obtain consistent data series that go back far enough to include a lot of temporal variation, and where random effects models are inappropriate because of correlation with the observed explanatory variables (as in this case).[4] These region and time fixed effects are included to capture other factors, e.g. seasonal factors, benefit generosity and screening intensity, not included in our data. Of course different British regions also inherited different starting positions in terms of IB rolls and the other observed factors in 1998Q1 and the regional fixed effects will be indistinguishable from differences in the temporal means of the observed variables. Consequently, our focus is on how within-region changes in these observed factors over our study period affect regional IB rolls, and we report the associated within R^2s rather than overall R^2s.

<Table 1 around here>

Correlations between observed variables are reported in Table 2. As we would expect, regional IB rolls are positively correlated with regional disability incidence and with unemployment rates, but negatively correlated with average earnings.

<Table 2 around here>

4. The Determinants of Male IB Stocks, Treated as Stationary

Perhaps the first decision to make in modelling the determinants of IB stocks over time, or using aggregate panel data as in this case, is whether they should be treated as stationary or non-

[4] Cross section approaches at a more disaggregated spatial level (e.g. Nolan and Fitzroy, 2003; McVicar, 2007) do not face the same constraint, but can tell us little about dynamics and face a greater risk of omitted variable bias because they do cannot control for unobserved heterogeneity across space.

stationary (e.g. see the earlier discussion of Disney and Webb, 1991). There will certainly be a high degree of persistence in the quarterly series of the number of IB claimants because the average duration of claim is around six years (see Anyadike-Danes & McVicar, 2007). A similar argument will apply to the number of people reporting themselves disabled. If we misspecify these series as I(0) when in fact they behave like I(1) series, we risk mistaking spurious regression for a genuine causal relationship. On the other hand, two arguments suggest we might wish to proceed under the assumption that these series are stationary. First, we express these variables as proportions of the working age populations, which, because they are bounded, implies they cannot be random walks. Second, we have just 33 quarters of data which is a little on the low side for high-powered time series analysis relying on temporal asymptotics.

In the end this comes down to an empirical question – do the series behave like they have unit roots or not – and as such we can use panel unit root tests to help provide guidance. Table 3 reports results from the t-test suggested by Im et al. (2003), which rejects unit roots for all series except the IB rolls, the proportion of the population aged over 50 and, at 95% but not 99%, the disability prevalence variable. Such tests have low power and other problems, however, so this is not strong evidence for or against unit roots. Because of this uncertainty and because of the nature of the data set, we explore estimation under both alternatives, first assuming stationarity and, in Section 5, assuming non-stationary like behaviour.

We begin by estimating the reduced form Equation (1) for the proportion of a region's working age male population claiming IB in quarter *t*. The included explanatory factors are the proportion reporting a disability, denoted Dis_{it}; the probability of getting a job given participation in the labour market as proxied by the regional unemployment rate, U_{it} (we also explore the non-disabled employment rate as an alternative measure); and, because IB is paid at a national rate, the value of IB relative to expected wages as proxied by median regional real full time weekly earnings, W_{it} (again we explore tenth percentile earnings as an alternative measure). The probability of receiving IB given a claim will depend on an individual's disability status – captured here in aggregate by Dis_{it} – and the intensity of medical screening, which is unobserved but arguably invariant across regions and therefore captured by time fixed effects. Regional fixed effects, v_i, capture time-invariant unobserved heterogeneity (and inevitably temporal means) across regions. Because many of those on the IB register in any given quarter will have entered the register in the past we include lags of the right hand side variables, but these are restricted to the first four lags then the eight and twelfth lags because of limited observations.

The equation we go on to estimate by fixed effects is therefore as follows:

$$(1)\quad \log IB_{it} = \sum_{t-\tau} \beta_{1\tau} \log Dis_{it} + \sum_{t-\tau} \beta_{2\tau} \log U_{it} + \sum_{t-\tau} \beta_{3\tau} \log W_{it} + v_i + v_t + \varepsilon_{it}$$

where IB_{it} denotes the proportion of the working age population claiming IB in area *i* at time *t*. Results are presented in column 2 of Table 4 and labelled 'FE1'. The within R^2 is 0.8, although this includes the explanatory power of the (unreported) time fixed effects. Neither do we report the estimated regional fixed effects, but note that they follow a 'north'/'south' pattern as might be expected from Figure 1.

According to the FE1 results, the proportion of the working age male population reporting a disability does not affect the male IB roll, either contemporaneously or with any of the included lags (they are jointly insignificant). These results are consistent with Faggio and Nickell's (2005) finding that self-reported disability incidence from the General Household Survey (GHS) has no significant impact on prime age male inactivity, and might perhaps lend some support to those aggregate panel studies that have omitted disability prevalence altogether (e.g. Disney and Webb, 1991). Note the conditioning on the fixed effects is likely to be important here – and similarly for Faggio and Nickell (2005) – given cross sectional evidence for Britain that suggests self reported disability prevalence *is* significantly correlated with IB rolls at the individual and LA level (e.g. Molho, 1989, McVicar, 2007). The different sets of results are reconcilable if there is substantial variation in disability prevalence across regions (captured here by the fixed effects) but not much variation across time within regions. For our data, the former is certainly true, although there is also *some* variation in disability prevalence over time within regions. (An additional factor to keep in mind is that measurement error due to different self-perceptions of disability may downward bias the estimated impact of disability prevalence in column 2 of Table 4 (see Bound, 1991).)

All but the twelfth lag of median weekly earnings is statistically significant at least at the ten percent level, and all are negatively signed. The suggestion is, consistent with the findings of Black et al. (2002), Autor and Duggan (2003) and Faggio and Nickell (2005), that economic incentives – what disability benefit claimants can potentially obtain in the local labour market – play a role in determining IB rolls. Because IB is paid at a national rate, the negative relationship with average earnings implies that the higher the replacement rate, the higher the IB roll. Summing the significant lags suggests a medium run elasticity of male IB rolls to regional average earnings of around -0.5, smaller than that suggested by Faggio and Nickell (2005) but closer to that suggested by Disney Webb (1991) for the UK and Black et al. (2002) for the US. This effect disappears when median earnings are replaced by earnings at the tenth percentile, which may be a more relevant measure for disabled workers in the lower part of the wage distribution. This may reflect the quality of the tenth percentile earnings data, which are employer-based rather than residence based, or it may reflect that the disabled really do react to median earnings rather than earnings lower down the distribution.

Unemployment rates are positively related to male IB rolls, with all lags statistically significant and with an overall medium run elasticity of 0.3. Given the concept of hidden unemployment we might want to explore alternative measures of the state of the labour market, e.g. employment rates. We check employment rates for the non-disabled (although these are not gender specific), in which case the labour market effects are correctly signed but no longer statistically significant.

Of course the results discussed above are for a somewhat arbitrary model in terms of the lags included and excluded (given the data constraints there was no testing down from a general model to the specific model discussed above). Klerman and Haider (2004) argue that this ad hoc approach to dynamics is a feature of much of the empirical welfare caseload literature and suggest analysis of flows rather than stocks as the best way forward. We follow their suggestion in Section 6, but for now explore inclusion of an LDV in a version of Equation (1) as an alternative to an arbitrary number of lags of explanatory variables. Dropping the multiple lags of the explanatory variables also allows us to more comfortably include a further two (admittedly blunt) controls for demographics and industry mix, specifically for the proportion of the working age population aged 50 years or over, Old_{it}, and the proportion of jobs that are in manufacturing, M_{it}. We therefore estimate the following LDV model, first by least squares dummy variables (LSDV) and then by Arrelano-Bond Generalized Method of Moments methods:

$$\log IB_{it} = \beta_0 \log IB_{it-1} + \beta_1 \log Dis_{it} + \beta_2 \log U_{it} + \beta_3 \log W_{it} + \beta_4 \log Old_{it} + \beta_5 \log M_{it} + v_i + v_t + \varepsilon_{it} \quad (2)$$

Results are presented in column 3 (FE2) of Table 4. The Arrelano-Bond results are identical to the LSDV results in terms of estimated coefficients, although there is a small difference in standard errors, so only the LSDV results are reported. Again, both the median earnings and the unemployment rate appear statistically significant and take the expected signs, with long run elasticities of 0.7 and -1.5 respectively. Also positive and statistically significant is the manufacturing share of employment. Disability prevalence is insignificant. Replacing the unemployment rate with the non-disabled employment rate gives the correct sign but not statistical significance; replacing median earnings with tenth percentile earnings gives the correct sign and retains statistically significance, albeit only at the ten percent level.

The key result from the LDV models, however, is that the coefficient on the LDV appears equal to one, i.e. it looks like the proportion of working age men claiming IB has a unit root over the sample period. So, we now have further cause to question the assumption of trend-stationarity for IB rolls. In Section 5 we start from the assumption of non-stationary series and consider cointegrating relationships and ECMs for IB rolls. Before moving on however, we check robustness by re-estimating the LDV models using annual data rather than quarterly data, taking the winter quarters for each year with quarterly values replaced by annual averages. Results are presented in column 4 of Table 4 (FE3). The quarterly and annual results are similar, i.e. apparent unit roots or close to unit roots and significant and correctly signed coefficients on median earnings, unemployment rates and manufacturing share.

According to these models, are changes in the observed variables enough to account for the ten percent fall in the national male IB roll over the sample period, and do they help explain the observed regional convergence in regional IB rolls? The statistically significant determinants from the FE1 estimates are unemployment rates and median earnings, both of which have moved in the right direction to explain the national fall. With respective falls of 16 percent and 12 percent, and associated elasticities (summing the significant lags) of 0.3 and -0.5, these changes would predict a fall in male IB rolls over the period of 11 percent, i.e. very close to the observed change. According to the FE2 estimates, the observed changes in unemployment rates and median earnings will *eventually* lead to a 19 percent fall in the national male IB stock, implying that the impact of falling unemployment rates and rising real earnings have not yet fully worked through into IB rolls. Both sets of estimates also predict regional convergence in IB rolls, driven primarily by convergence in regional unemployment rates.

5. The Determinants of Male IB Stocks, Treated as Non-stationary

Given the a priori doubt over whether IB rolls could be treated as stationary, together with the evidence of unit roots from the IPS tests presented in Table 3 and from the LDV models presented in Table 4, this section explores the determinants of IB rolls treating them explicitly as non-stationary. Our interest lies in both the long run determinants of IB rolls and the short run dynamics about the equilibrium relationship, should one exist. In other words, we are interested first in the existence or otherwise of one or more cointegrating relationships between the proportion of working age men claiming IB, disability prevalence, labour market and other factors; second in the nature of any such equilibrium relationship(s); and third in the ECM governing the dynamics around any such equilibrium relationship.

Although there are methods for testing for cointegrating relationships in panel data settings which could in principle be applied here (e.g. see Pedroni, 1999), such methods may be of limited use given the length of our sample period. For convenience, we therefore *assume* that a long run equilibrium relationship exists and proceed to estimate an ECM. Results are presented in Table 5 for both quarterly and annual data versions of the model. The discussion concentrates on estimation of unrestricted ECMs (ECM2 and ECM4 in the Table), but we also present results for ECMs with an imposed long run relationship (ECM1 and ECM3 in the Table).[5]

Specifically, we estimate the following equation:

$$(3)\quad \begin{aligned}&\Delta\log IB_{it} = \Delta\beta_1 \log Dis_{it} + \Delta\beta_2 \log U_{it} + \Delta\beta_3 \log W_{it} + \Delta\beta_4 \log Old_{it} + \Delta\beta_5 \log M_{it} + \ldots\\ &\ldots\lambda(\log IB_{it-1} - \gamma_1 \log Dis_{it-1} - \gamma_2 \log U_{it-1} - \gamma_3 \log W_{it-1} - \gamma_4 \log Old_{it-1} - \gamma_5 \log M_{it-1}) + v_i + v_t + \varepsilon_{it}\end{aligned}$$

For the ECM2 and ECM4 models, the γs are recoverable by dividing the estimated coefficient on the lag level terms by (minus) the estimated coefficient on the lag level dependent variable.

[5] The ECM1 and ECM3 models impose the following restrictions based on the cross section results from McVicar (2007): γ_1=1, γ_2=0.5, γ_3=-1, γ_4=0 and γ_5=0.

As we might expect given average claim duration of six years, the ECM2 and ECM4 results suggest slow convergence to the long run equilibrium following a shock. There are correctly signed impacts on IB rolls from contemporaneous changes in median earnings and unemployment rates and also from contemporaneous changes in the share of employment in manufacturing, although this is only statistically significant for the quarterly data. The directly estimated cointegrating vector is not unlike that imposed for ECM1 and ECM3, with similar coefficients on median earnings (γ_3=-1.2), and on unemployment rates (γ_2=0.6), although there is no long run relationship evident between IB rolls and disability prevalence in the directly estimated models.

These results change little if we substitute non-disabled employment rates for unemployment rates (opposite sign as we expect), or tenth percentile earnings for median earnings (same sign but not as consistently significant as median earnings). Our conclusion from this exercise is therefore that labour market factors – earnings and unemployment rates – are again the most consistent determinants of IB rolls over the period; disability prevalence again does not appear to have a significant relationship with IB rolls either in the short run or the long run, at least over this period. Given observed changes in the labour market at national level over this period, the ECM2 estimates predict an eventual fall in the male IB stock of 23 percent, again implying there are further falls in male IB rolls to come. Similarly, the model predicts convergence in regional male IB rolls over the period given observed convergence in regional unemployment rates.

6. The Determinants of IB Flows

Klerman and Haider (2004) argue that analysis of flows rather than stocks is the best way forward for modelling welfare caseloads for precisely the kinds of reasons we have discussed in the previous sections. IB flow data are only publicly available for British regions since 1999Q3, however, and are described by the ONS as 'experimental'. Nevertheless we estimate the following expressions for the outflow rate and inflow rate, again using both quarterly and annual data, allowing for two quarterly lags or one annual lag:[6]

$$(4)\quad In_{it} = \sum_{t-\tau} \beta_{1\tau} Dis_{it} + \sum_{t-\tau} \beta_{2\tau} U_{it} + \sum_{t-\tau} \beta_{3\tau} W_{it} + \sum_{t-\tau} \beta_{4\tau} Old_{it} + \sum_{t-\tau} \beta_{5\tau} M_{it} + \nu_i + \nu_t + \varepsilon_{it}$$

$$(5)\quad Out_{it} = \sum_{t-\tau} \beta_{1\tau} Dis_{it} + \sum_{t-\tau} \beta_{2\tau} U_{it} + \sum_{t-\tau} \beta_{3\tau} W_{it} + \sum_{t-\tau} \beta_{4\tau} Old_{it} + \sum_{t-\tau} \beta_{5\tau} M_{it} + \varphi_i + \varphi_t + \psi_{it}$$

Results are presented in Table 6. First consider quarterly inflows. Disability prevalence and its lags are statistically insignificant. Inflows do appear, however, to be significantly related to median earnings (albeit only the second lag and only at the ten percent level) and unemployment rates. A one pound increase in median earnings leads to a .000003 percentage point fall in inflows to IB, which given a sample mean of .007 corresponds to an elasticity of -0.2. A one percentage point increase in the unemployment rate leads to an immediate .00017 percentage point increase in IB inflows and a further two such increases in the following two quarters, with the overall elasticity of inflows to the unemployment rate around 0.4. A one percentage point increase in the share of employment in manufacturing leads to a -.0003 percentage point increase

[6] Even with the flow data, at least for inflows, we need to account for possible dynamics. because many claimants go through a six month period on statutory sick pay (SSP) before showing up on the IB register.

in IB inflows, corresponding to an elasticity of -0.6. Note in each case the significance of the second lag terms is consistent with the effects of SSP, i.e. the six month delay between leaving employment on health or disability grounds and eligibility for IB. The annual inflow estimates suggest similarly signed and broadly similarly sized impacts, although only the unemployment rate terms retain statistical significance together with a now significant positive impact of the population share over 50 with elasticity 0.25.

Turning to the outflows equation, again there is a marginally significant impact of lagged median earnings, with elasticity of 0.1, and of lagged unemployment rates with elasticity -.05. Additionally, however, there is an intuitively signed but only marginally significant impact of lagged disability prevalence on outflows with an elasticity of -0.15, and – the strongest determinant – a significant positive impact of the population share over 50 with an elasticity of 1.4. This latter effect suggests the population share over 50 may be capturing outflows from IB to pensions. Only the unemployment rate and population share over 50 are significant in the annual version of the model, with the same signs.

We can use the estimated flow effects to simulate what would happen to the IB stock over time as a result of changes in the determinants of the flows. Here we use the estimates for the quarterly flow models from Table 6, with statistically insignificant coefficients set to zero. First consider a country-wide 'shock' to disability prevalence, say a 50 percent increase (rather large but just to more clearly see the effect). Through its impact on IB outflows this would lead to an increase of six percent in the IB stock over a period of 32 quarters. Second consider a similar sized shock to the unemployment rate, i.e. a 50 percent increase. Acting mainly through inflows this would lead to a 22 percent increase in the IB stock over a similar period. (As before, regional convergence in unemployment rates creates pressure for convergence in regional IB rolls.) A similarly sized shock to median earnings, acting through both inflows and outflows, would lead to a 13 percent fall in the national IB stock over 32 quarters. A 50 percent increase in the population share aged over 50 would lead to a 40 percent fall in the IB stock. Finally, a 50 percent increase in the manufacturing share of employment would lead to a 27 percent fall in the IB stock.

In fact, nationally over the sample period there has been a ten percent increase in disability prevalence, a 16 percent fall in unemployment rates, a twelve percent increase in median earnings, a seven percent increase in the share of the working age population aged over 50, and a 30 percent fall in the manufacturing share of employment. Our flows model suggests that these changes would have together led to a three percent fall in the overall IB stock, with positive labour market and demographic changes offset by increased disability prevalence and structural change. The data actually show a ten percent fall, so changes in observed variables at the national level appear insufficient to explain the full extent of the fall in the national male IB roll when using the flows model. The explanation for this is that most of the rest of the observed fall in the national IB stock over the period is ‘driven’ in the flows model by the time dummies in the outflows equation. In the absence of major administrative changes to IB, these time dummies are most likely capturing the effects of changes to the age profile of IB claimants over the sample period, which has seen a disproportionately large number of ageing claimants move off IB into retirement.[7]

[7] Claimant data by age group show a sharp fall in the proportion of male IB claimants that are aged 60-64 since 1999Q3.

7. Summary and Conclusions

This paper provides quantitative evidence at the UK regional level on commonly suggested determinants of male IB rolls over the period 1998-2006, adopting three alternative approaches for dealing with dynamics. Conditional on time and region fixed effects, self-reported disability prevalence appears largely unrelated to IB rolls over this period. Rather the evidence suggests that male IB claimants in Britain are responding to economic incentives in terms of labour market opportunities available to them. Specifically, changes in unemployment rates and median earnings explain part of the fall in the national male IB roll over the sample period.[8] According to the three models, observed changes in these factors over the sample period have led to falls of between three and 11 percent in the proportion of the working age male population claiming IB over this period (compared to an observed fall of ten percent nationally), with *eventual* falls predicted of up to 23 percent once all long run effects have worked through. Regional convergence in unemployment rates over the period also goes some way to explaining the observed convergence in regional IB rolls over the period.

The paper makes a number of contributions. First, it adds evidence on the determinants of British IB rolls, using recent data, to a literature that contains little in the way of existing quantitative evidence, with what there is largely dating back to the late 1980s or early 1990s. Second, the paper includes (self-reported) disability prevalence in a literature where it is often omitted as a potential determinant of IB stocks, although it appears insignificant over this period. Third, the

[8] Note the contrast with the US where evidence suggests that DI stocks have risen as a result of falling real earnings and rising replacement rates.

paper pays far more attention to the dynamics of IB stocks than existing aggregate level British studies and some US studies. Finally, the paper highlights the limitations inherent in such an exercise when using aggregate data, and short series of aggregate data at that.

The British government has recently set a target of reducing IB rolls by one million over the next ten years (see Freud, 2007). A variety of measures have been put in place towards this aim, including the strengthening of the link between receipt of IB and job search that forms the key part of Pathways to Work, as well as various changes to IB rates and eligibility rules. Tougher medical screening is also on the way with expected reforms to the Personal Capability Assessment in 2008. Taken together these reforms are likely to further reduce IB rolls, and early evidence on Pathways suggests just that, although the impact to date has not been large (see Blyth, 2006).

What this paper shows, however, is that the economic environment is crucial in explaining IB claims, at least for males. As unemployment rates have fallen over recent years, so too have male IB rolls. Indeed, the evidence in this paper suggests that the effects of falling unemployment may not yet have fully worked through into male IB stocks, so that there are further falls to come. On the down side, however, further significant falls in unemployment are unlikely, so policy makers may have to rely on falling replacement rates if they are to meet the one million target, particularly if the labour market takes a turn for the worse. This sits less comfortably with government's obligation to provide a reasonable standard of living to the work limiting disabled population. Tougher screening might also contribute – although the evidence here suggests that recent changes in IB stocks are not closely related to changes in the number of men reporting themselves disabled – but there are limits to how far down this road we can go if we are not to deny incapacity benefits to those who really need them.

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Figure 1: Regional IB Rolls, Male working age, ratio to population, 1998Q1 & 2006Q1

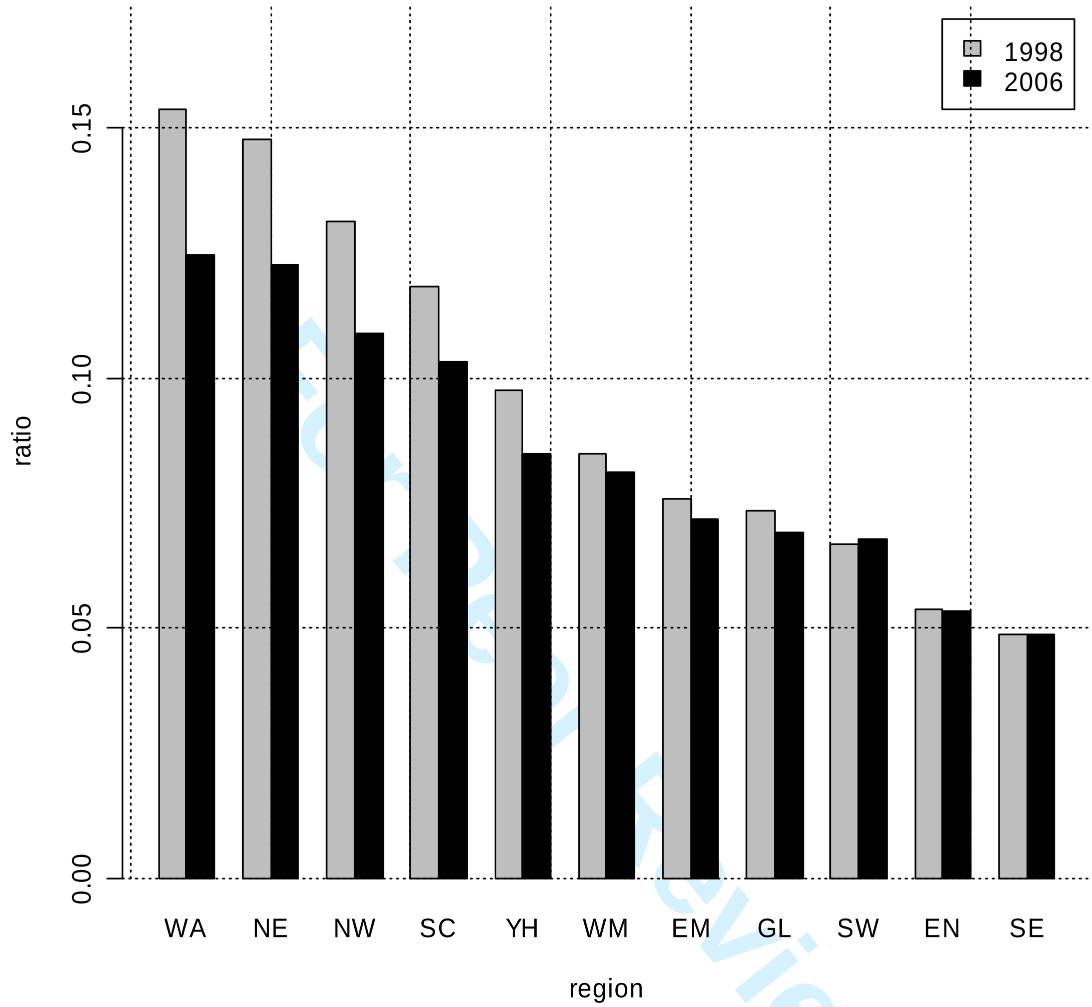
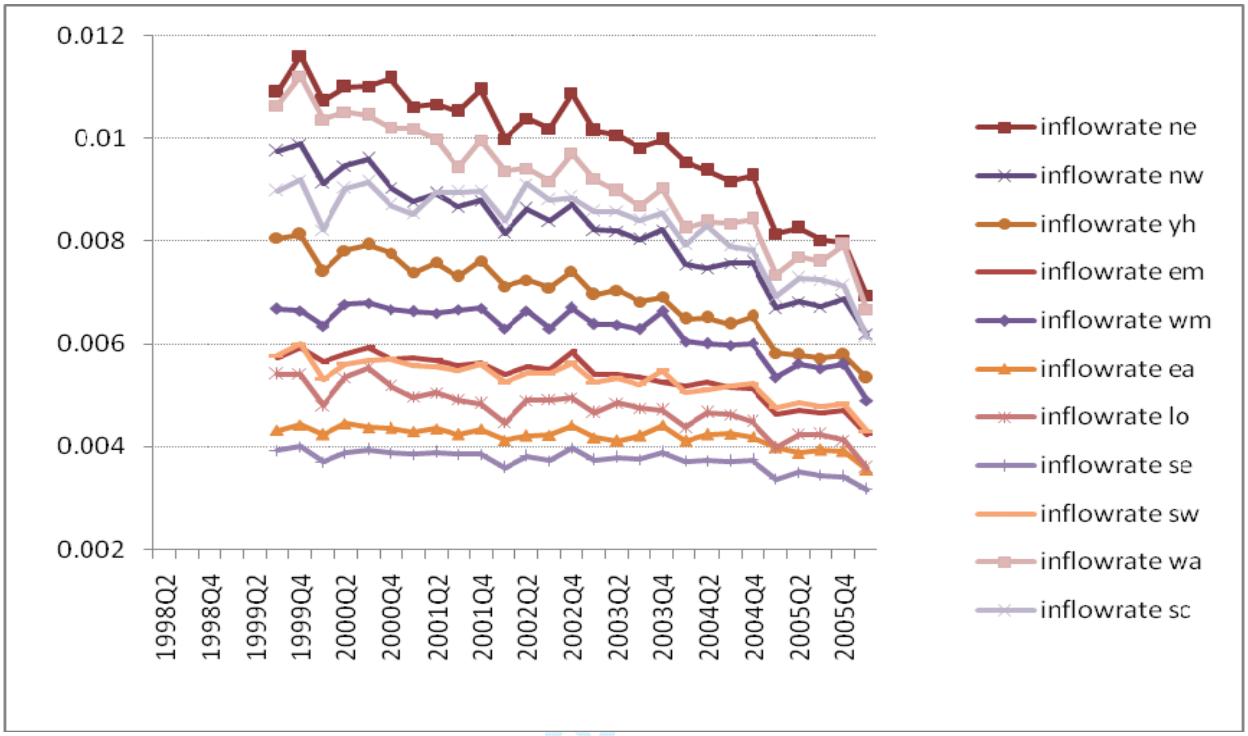
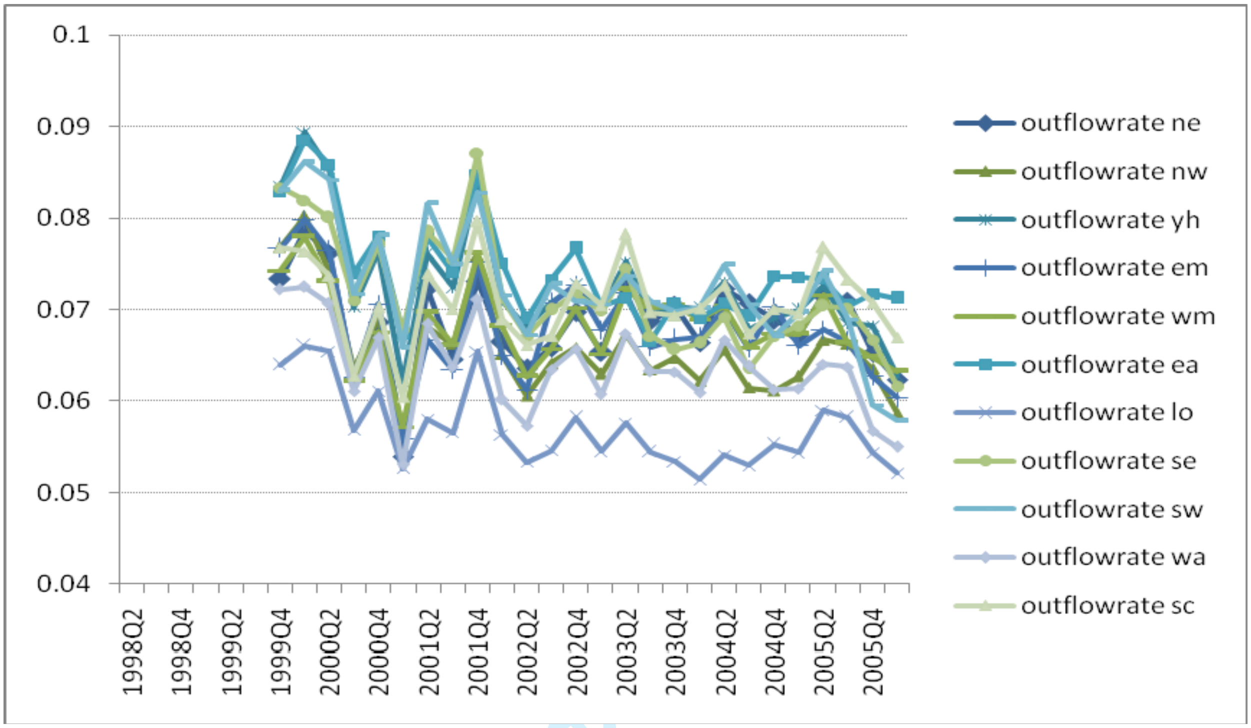


Figure 2: Male IB Inflow Rates, 1999Q3-2006Q1, Government Office Regions



Note: The vertical axis shows the proportion of the at-risk population (working age men not claiming IB) starting a claim in each quarter

Figure 3: Male IB Outflow Rates, 1999Q4-2006Q1, Government Office Regions



Note: The vertical axis shows the proportion of the IB stock ending a claim over the quarter.

Table 1: Sample Means and Standard Deviations, by Region, 1998Q1-2006Q1

	IB Claimants/Working Age Male Pop	Male Working Age Disability Prevalence	Median Real Weekly Earnings, Full-Time, £	10th Percentile Real Weekly Earnings, Full-Time, £	Unemployment Rate	Population Share Over 50 Years Old (Males and Females)	Manufacturing Share of Employment (Males and Females)
Northeast	.136 (.006)	.254 (.015)	351 (12.7)	196 (5.61)	.086 (.015)	.250 (.007)	.163 (.024)
Northwest	.119 (.005)	.212 (.008)	372 (15.3)	203 (6.82)	.061 (.011)	.250 (.004)	.162 (.022)
Yorks & Humber	.091 (.003)	.208 (.008)	367 (13.4)	204 (7.33)	.063 (.011)	.248 (.005)	.172 (.021)
East Midlands	.074 (.001)	.187 (.010)	378 (15.9)	204 (8.52)	.051 (.005)	.255 (.006)	.203 (.025)
West Midlands	.083 (.002)	.200 (.012)	376 (14.4)	210 (7.13)	.063 (.007)	.253 (.005)	.201 (.030)
Eastern	.053 (.001)	.166 (.009)	445 (18.3)	216 (8.21)	.042 (.005)	.255 (.005)	.141 (.016)
London	.071 (.001)	.161 (.006)	518 (22.5)	252 (10.7)	.077 (.006)	.190 (.003)	.065 (.009)
Southeast	.048 (.001)	.162 (.008)	472 (15.4)	226 (9.90)	.039 (.004)	.251 (.004)	.111 (.014)
Southwest	.067 (.001)	.189 (.008)	389 (14.2)	204 (8.03)	.042 (.006)	.270 (.006)	.139 (.017)
Wales	.138 (.007)	.240 (.009)	358 (17.0)	196 (8.08)	.066 (.013)	.265 (.006)	.175 (.019)
Scotland	.111 (.003)	.202 (.007)	384 (15.3)	204 (6.21)	.073 (.010)	.244 (.007)	.124 (.020)
Britain	.090 (.030)	.200 (.030)	401 (53.4)	210 (17.5)	.060 (.017)	.249 (.020)	.150 (.044)

Note: Standard deviations in parentheses.

Table 2: Pairwise Correlations between Observed Variables

	IB Claimants/Working Age Male Pop	Male Working Age Disability Prevalence	Median Real Weekly Earnings, Full-Time, £	10th Percentile Real Weekly Earnings, Full-Time, £	Unemployment Rate	Population Share Over 50 Years Old (Males and Females)
Disability Prevalence, Standard Definition, Male	.871					
Median Real Weekly Earnings	-.660	-.713				
10th Percentile Real Weekly Earnings	-.541	-.603	.938			
Unemployment Rate	.662	.429	-.266	-.131		
Population Share Over 50 (All)	.123	.386	-.584	-.648	-.479	
Manufacturing Share of Employment (All)	.317	.392	-.797	-.737	.116	.485

Table 3: IPS Panel Unit Root Tests, T-test Statistics

Variable (logs)	T-test
IB Claimants/Working Age Pop	-1.23
Disability Prevalence (Standard Definition)	-2.74**
Median Real Weekly Earnings	-3.96***
Unemployment Rate	-3.56***
Population Share Over 50	-1.59
Manufacturing Share of Employment	-2.65**

Notes: *** (**) denote rejection of the null (of I(1)) at 1% (5%) levels.

Table 4: IB Stock Models Assuming Stationarity

	FE1 (quarterly)	FE2 (quarterly)	FE3 (annual)
Disability Prevalence	-.044 (.030)	-.003 (.005)	-.003 (.033)
...L(1)	.009 (.023)		
...L(2)	.011 (.020)		
...L(3)	.004 (.019)		
...L(4)	-.028 (.030)		
...L(8)	-.054 (.032)		
...L(12)	-.012 (.044)		
Median Real Weekly Earnings	-.056 (.041)	-.021** (.007)	-.237** (.076)
...L(1)	-.087* (.046)		
...L(2)	-.119** (.052)		
...L(3)	-.133** (.049)		
...L(4)	-.099** (.036)		
...L(8)	-.092** (.039)		
...L(12)	-.033 (.022)		
Unemployment Rate	.039*** (.009)	.010** (.004)	.051*** (.015)
...L(1)	.026** (.008)		
...L(2)	.039*** (.007)		
...L(3)	.022** (.009)		
...L(4)	.046*** (.006)		
...L(8)	.061*** (.016)		
...L(12)	.038** (.014)		
Population Share Over 50		.037 (.040)	.050 (.193)
Manufacturing Share of Employment		.026** (.009)	.101** (.040)
Time Dummies	Yes***	Yes***	Yes**
IB Claimants/Working Age Pop (L1)		.986*** (.008)	.920*** (.048)
R^2 (within)	.797	.988	.960
Observations	220	341	77

Notes: Robust standard errors in parentheses. *** (**, *) statistically significant at1% (5%, 10%). All variables are in logs. The dependent variable is the log of the proportion of the working age male population claiming IB.

Table 5: IB Stocks Assuming Non-Stationarity

	Quarterly Data		Annual Data	
	ECM1	ECM2	ECM3	ECM4
Δ Disability Prevalence	-.005 (.005)	-.008 (.006)	.039 (.030)	-.026 (.034)
Δ Median Real Weekly Earnings	-.010 (.006)	-.020** (.008)	-.116** (.045)	-.209** (.081)
Δ Unemployment Rate	.002 (.002)	.008** (.003)	.029* (.015)	.045*** (.012)
Δ Population Share Over 50	-.036** (.016)	-.004 (.046)	-.081 (.125)	-.059 (.150)
Δ Manufacturing Share of Employment	.017 (.015)	.037** (.014)	.052 (.084)	.113 (.070)
IB Proportion L(1)	-.006* (.003)	-.024** (.010)	-.077* (.036)	-.175** (.064)
Disability Prevalence L(1)		-.001 (.005)		-.022 (.030)
Median Real Weekly Earnings L(1)		-.029** (.011)		-.333** (.133)
Unemployment Rate L(1)		.014*** (.003)		.085*** (.018)
Population Share Over 50 L(1)		.036 (.041)		.031 (.163)
Manufacturing Share of Employment L(1)		.025** (.009)		.116** (.047)
Time Dummies	Yes***	Yes***	Yes*	Yes***
R^2 (within)	.803	.823	.865	.921
Observations	341	341	77	77

Notes: Robust standard errors in parentheses. *** (**, *) statistically significant at1% (5%, 10%). All variables are in logs. The dependent variable is the first difference of the log of the proportion of the working age male population claiming IB.

Table 6: Male IB Outflow & Inflow Rates

	Outflows		Inflows	
	Outflows 1 (quarterly FE)	Outflows 2 (annual FE)	Inflows 1 (quarterly FE)	Inflows 2 (annual FE
Disability Prevalence	.005 (.030)	.019 (.048)	.0006 (.002)	-.005 (.005)
...L(1)	.018 (.025)	.028 (.039)	-.0006 (.003)	.0003 (.003)
...L(2)	-.048* (.024)		-.004 (.004)	
Median Real Weekly Earnings	.00002 (.00002)	-.00003 (.00002)	-.000001 (.000002)	-.000002 (.000001)
...L(1)	.00002* (.00001)	-.00001 (.00001)	-.000004 (.000003)	-.0000001 (.000002)
...L(2)	.00002 (.00001)		-.000003* (.000002)	
Unemployment Rate	-.017 (.035)	-.007 (.023)	.017*** (.004)	.006** (.002)
...L(1)	-.017 (.030)	-.038*** (.010)	.012** (.005)	.010*** (.001)
...L(2)	-.061** (.027)		.019*** (.005)	
Population Share Over 50	.398** (.153)	.208*** (.061)	-.019 (.019)	-.005 (.005)
...L(1)	.106 (.179)	-.070 (.058)	.010 (.011)	.007** (.003)
...L(2)	-.093 (.149)		.016 (.022)	
Manufacturing Share of Employment	.013 (.086)	-.018 (.022)	.012 (.016)	-.001 (.005)
...L(1)	-.050 (.056)	-.015 (.032)	.015 (.014)	.002 (.004)
...L(2)	-.121 (.086)		-.027** (.008)	
Time Dummies	Yes***	Yes***	Yes***	Yes***
R^2 (within)	.862	.588	.852	.943
Observations	264	66	275	77

Notes: Robust standard errors in parentheses. *** (**, *) statistically significant at1% (5%, 10%). Coefficients can be read as percentage point impacts of a one unit change in each case (the dependent variables are the number of outflows divided by the lagged stock and the number of inflows divided by the lagged at risk (i.e. non-claiming) population).